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Inflation and Interest Rate Differentials Between Germany and its EMS Partners*

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1. Introduction

Most observers agree that the European Exchange Rate Mechanism (ERM) has been a major factor in prompting its inflation-prone member countries to disinflate. In a series of papers *Giavazzi* and *Giovannini* (1989, 1985) forcefully argued that the ERM is best described as an asymmetric exchange rate system. Germany as the center country of this system sets its monetary policy independent of other ERM members who are bound to accommodate the German policy stance or, if unwilling to do so, see their exchange rates depreciate relative to the Deutsche Mark (DM). Expressing this notion, the ERM has frequently been called a "greater Deutschmark area" (*Herz* and *Röger*, 1992; *Giavazzi* and *Giovannini*, 1987).

Recently, several contributions have approached the issue of asymmetries empirically. The majority of the empirical work suggests that the monetary policy adopted by the German Bundesbank plays a predominant role in the European Monetary System (EMS) (*Funke* and *Hall,* 1994; *Kirchgässner* and *Wolters,* 1993; *von Hagen* and *Fratianni,* 1990). Hoping to borrow monetary credibility from the Bundesbank, countries with higher inflation rates than Germany pegged the respective exchange rate to the Deutsche Mark (*Giavazzi* and *Pagano*, 1988).[1] Persistent inflation differentials in combination with fixed exchange rates, however, led to large swings in competitiveness across Europe. These were at least partially offset by a series of realignments between 1980 and 1987 but the 1992 EMS turbulence, following a period of apparently stable parities, virtually abandoned the concept of a comprehensive fixed exchange rate regime with temporary realignments.

In the present paper we address the expectations of market participants on EMS realignments. Inflation differentials cause market participants to

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1 For a critical comment see *Obstfeld* (1988).

entertain expectations on corrective exchange rate adjustments. While a floating exchange rate regime allows for such an adjustment to occur instantaneously, a fixed exchange rate system may in the short-run prevent the expected exchange rate adjustments to materialize. Anticipating the impending realignment, rational market participants will therefore demand a compensating return on foreign-currency assets which should cause the emergence of a new, or the widening of an existing interest rate differential. We suppose long-term interest rate differentials between Germany and its EMS partners to serve as a proxy for exchange rate expectations and examine their link with inflation differentials. If market participants are suspicious about the sustainability of a narrowed inflation differential, it may be expected that for a country with lower credibility a decrease in the inflation differential entails a smaller effect on interest rate differentials than a similar decrease for a country with higher credibility.

Subsequently we consider three different aspects of the relationship between interest rate and inflation differentials. First, both interest rate and inflation differentials appear to be non-stationary for the majority of the countries under study. However, if realignments are intended to offset changes in competitiveness, one should find a stationary long-run relationship between the two. Second, by affecting market expectations on future realignments, inflation differentials should Granger-cause interest rate differentials. Third, asymmetries within the EMS suggest an asymmetric response of realignment expectations to deviations from the equilibrium exchange rate. If inflation differentials vis-à-vis Germany have driven the exchange rate of an EMS member country currency below its equilibrium level, market participants would nevertheless expect a nominal appreciation to be unlikely. This notion translates into a slower adjustment of interest rate differentials to long-run equilibrium in response to a negative deviation. Observations pertaining to the 1990-1993 period back these considerations. In the aftermath of unification, Germany found itself in the company of countries that were traditionally underperforming in terms of price level stability such as Italy, whereas France and Denmark performed remarkably better. Despite the narrowing and even the reversion of the inflation differential, the nominal interest rate differentials to Germany were, at best, eliminated but did not turn negative. The spread between ex-post real interest rates across countries thus widened.

The paper is organized as follows. The subsequent chapter 2 discusses the theoretical link between inflation differentials and interest rate differentials. In particular, we draw upon the concepts of uncovered interest parity and purchasing power parity to derive the principal equation of our econometric model. Chapter 3 introduces the econometric model. We conduct a cointegration and Granger causality analysis for long-term interest rate and inflation differentials of EMS members vis-à-vis Germany using a bivariate vector autoregression (VAR) framework. The empirical results are presented in chapter 4. Chapter 5 concludes the paper.

2. Inflation and Interest Rate Differentials: A Theoretical Link

Uncovered interest parity (UIP) as well as real interest parity (RIP) are concepts which link financial markets in different countries.[2] Uncovered interest parity holds if the nominal interest rate of currency A, i_t^A, plus its expected appreciation (or depreciation, respectively) equals the nominal interest rate of currency B, i_t^B,

$$\text{(1)} \qquad \text{UIP}: \quad i_t^A - i_t^B = \mathrm{E_t}(S_{t+1} - S_t)$$

or

$$\text{(2)} \qquad \text{UIP}: \quad di_t = S_{t+1}^e - S_t$$

with $\mathrm{E_t}$ denoting the conditional expectations operator with respect to the information set available at period t, di_t the nominal interest rate differential in period t, S_t the spot rate of currency B in terms of currency A (in logs), and S_{t+1}^e the value of S_{t+1} expected as of period t. Short-term deviations from UIP (*nominal* interest parity) depend on three different factors. First, capital controls together with the risk of future controls – political risk according to *Aliber* (1973) – may lead to the emergence of a country premium on returns across national markets. Second, as the ex-post return of portfolio investments and of direct investments strongly depends on future exchange rates, the volatility of exchange rates may drive wedges between ex-ante returns in different markets and may therefore give rise to an exchange risk premium in excess of the expected appreciation (or depreciation) (*Marston,* 1993).[3] Post-tax yield differences constitute a third reason for deviations from nominal interest parity, and thus the average uncovered interest differential should equal the sum of country and exchange risk premiums as well as post-tax yield differences (*Isard*, 1991).

Since expected future exchange rates are not directly observable, in testing the empirical validity of UIP the most important strand of studies draws

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2 See *Hodrick* (1987) and *Cumby* and *Obstfeld (*1984) for reviews. An extensive individual study on the degree of financial market integration as measured by deviations from UIP has been provided by *Frankel* and *MacArthur (*1988).

3 In terms of the EMS before enlarging the bands the primary risk was that of future realignments.

upon the hypothesis that forward exchange rates should serve as unbiased predictors of future spot exchange rates (*Froot* and *Frankel*, 1989; *Fama*, 1984). Disregarding transaction costs, the absence of riskless arbitrage possibilities implies that

$$\text{(3)} \qquad CIP: \quad di_t = F_t - S_t$$

where F_t denotes the forward exchange rate.[4] Tests of the joint hypothesis that both this covered interest parity (CIP) and the unbiasedness hypothesis hold, however, have regularly been rejected (*MacDonald* and *Taylor*, 1989; *Meese*, 1989). Reexamining the uncovered interest parity relationship, *McCallum* (1992) confirms the earlier results on forward rate biasedness but demonstrates that this evidence does not necessarily entail rejection of UIP. If monetary authorities are assumed to manage interest rate differentials so as to avoid rapid changes in exchange rates and in the related differentials, the failure of unbiasedness is nevertheless consistent with UIP. The results of empirical studies on Euromarkets indeed suggest that deviations from uncovered interest parity are mainly due to a country premium associated with capital controls or political risk and post-tax yield differences (*Marston,* 1993; *Dooley* and *Isard*, 1980) and that UIP holds in the long run.

Long-run exchange rate expectations, in turn, are related to cross-country inflation differentials. In its weak form purchasing power parity (PPP) states that inflation differentials are equal to the change of the nominal exchange rate,

$$\text{(4)} \qquad \text{PPP}: \quad p^A_{t+1} - p^B_{t+1} = dp_{t+1} = S_{t+1} - S_t \ ,$$

with p^A (p^B) as the inflation rate of country A (B) and dp as the inflation differential between countries A and B. While tests of PPP for floating exchange rate regimes provide mixed evidence (*Juselius*, 1995; *MacDonald, 1995)*,[5] for a managed exchange rate regime like the EMS the validity of PPP in its weak form is subject to discretionary policy intervention.

Subtracting the expected inflation differential $\mathrm{E}_t(dp_{t+1})$ from both sides of equation (1), we obtain an expression for ex-ante *real* interest parity,

4 Empirical support for this relationship is surveyed in *MacDonald* (1988) and *Levich* (1985).

5 Tests of the strong form of purchasing power parity regularly fail to give evidence in favor of its validity. Several reasons have been advanced to explain this finding: Consumer price indices might differ due to the inclusion of non-traded goods or the application of different weighing schemes across countries *(Dornbusch,* 1985). A reason other than such measurement problems is differences in the elasticity of substitution in consumption across countries which allow firms to exert some degree of price discrimination. If a firm can at least partially control the cross-border trade in its good(s), the real exchange rate can deviate from one (*Tootell*, 1992).

(5) $$\text{RIP}: \quad dr_t = \mathrm{E}_t(S_{t+1} - S_t) - \mathrm{E}_t(dp_{t+1})\,.$$

If we assume that the risk of capital controls and post-tax yield differences is either negligible or constant over time, the hypothesized direct link between nominal interest rate differentials and inflation differentials may be caused by deviations from PPP, time-varying exchange risk premia or both (*Capitelli* and *Schlegel*, 1991). It seems reasonable to suppose that market participants expect cumulated inflation differentials in the long run to be offset by successive realignments which counteract the deterioration of competitiveness. Therefore time-varying exchange risk premia should have no effect on uncovered interest parity over long sample periods which is tantamount to stating that the right hand side of equation (5) equals zero. We thus have

(6) $$di_t - \mathrm{E}_t(dp_{t+1}) = 0$$

which represents the principal equation for our econometric model. Since PPP is a long-term concept, equation (6), as it stands, stipulates a long-run relationship between inflation and interest rate differentials. Deviations from PPP induce an adjustment process to equilibrium. To capture the associated short-run dynamics, we draw upon a dynamic model specification that will be introduced subsequently.

3. Methodology

Consider a 2-dimensional Vector Autoregressive (VAR) representation of order m for interest rate and inflation differentials (di_t, dp_t),

(7) $$\begin{pmatrix} di_t \\ dp_t \end{pmatrix} = \begin{pmatrix} \mu_1 \\ \mu_2 \end{pmatrix} + \begin{pmatrix} a_{11}(L) & a_{12}(L) \\ a_{21}(L) & a_{22}(L) \end{pmatrix} \begin{pmatrix} di_{t-1} \\ dp_{t-1} \end{pmatrix} + \begin{pmatrix} \epsilon_{1t} \\ \epsilon_{2t} \end{pmatrix}$$

where $\mu = (\mu_1, \mu_2)'$ is a vector of constant drift, and $\epsilon_t = (\epsilon_{1t}, \epsilon_{2t})'$ is white noise.[6] Furthermore it is assumed that the roots of the characteristic polynomial are strictly outside the unit circle or equal to one, $|\mathrm{I} - \mathrm{A}(z)| = 0 \Rightarrow |z| \geq 1$. Crucial for the behavior of the system is the number of unit roots in $\mathrm{A}(L)$. If there are no unit roots, that is, $|\mathrm{I} - \mathrm{A}(z)| = 0 \Rightarrow |z| > 1$, the VAR (7) is stable and (di_t, dp_t) follow a stationary process. If there are two unit roots, the system is nonstationary while in the intermediate case of one unit root the system can be rewritten in the vector error correction (VEC)

[6] In order to derive (7) from (6), it is necessary to assume static expectations, $\mathrm{E}_t(dp_{t+1}) = dp_t$.

representation with one error-correction term (see *Engle* and *Granger*, 1987),

$$(8)\quad \begin{pmatrix}\Delta di_t\\ \Delta dp_t\end{pmatrix} = \begin{pmatrix}\mu_1\\ \mu_2\end{pmatrix} + \begin{pmatrix}b_{11}(L)\ b_{12}(L)\\ b_{21}(L)\ b_{22}(L)\end{pmatrix}\begin{pmatrix}\Delta di_{t-1}\\ \Delta dp_{t-1}\end{pmatrix} + \begin{pmatrix}\alpha_1\\ \alpha_2\end{pmatrix}(di_{t-m} - \beta dp_{t-m}) + \begin{pmatrix}\epsilon_{1,t}\\ \epsilon_{2,t}\end{pmatrix}$$

where $b_{ij}(L)$ are lag-polynomials of order $m-2$. This is equivalent to the case that both dp_t and di_t, taken individually, have a unit root and are thus integrated processes while there exists a stationary linear combination $di_t - \beta dp_t$, that is the series are cointegrated. The adjustment coefficents α_i describe the speed of adjustment of the particular series to the long-run equilibrium while the lag polynomial $b(L)$ represents additional short-run dynamics. *Engle* and *Granger* (1987) proposed a two-step estimator of (8). The first step involves a regression of dp_t on di_t and a subsequent test for the stationarity of the residual. In the case of cointegration the estimated equilibrium error $di_t - \hat{\beta} dp_t$ is inserted into equation (8), and $b(L)$ and $\alpha = (\alpha_1, \alpha_2)'$ are estimated by OLS. This procedure is less efficient than the maximum likelihood estimator proposed by *Johansen* (1991).

In order to analyze the time-related interaction between the two series, we perform Granger-causality tests *(Granger,* 1969). Granger-causality from dp_t to di_t means that the conditional forecast for di_t can be significantly improved by adding lagged dp_t to the information set. This amounts to a non-zero polynomial $a_{12}(L)$ in the bivariate VAR (7). The feasibility of Granger causality tests depends on the stationarity features of the system. If the series are stationary, the null hypothesis of no causality can be tested by standard Wald tests (e.g. *Lütkepohl,* 1991). With integrated processes the situation generally becomes intractable as the distribution of the Wald statistic is nonstandard (*Toda* and *Philips,* 1993). However, for the case of a bivariate cointegrated system the standard distribution theory applies and conventional test procedures can be used. *Toda* and *Philips* (1991) have pointed out that the null hypothesis of dp_t not Granger-causing di_t is equivalent to $b_{12}(L) = 0$ and $\alpha_1 = 0$ in equation (8). They proposed to test each subhypothesis sequentially using the same significance level. Their simulations show that this method is in many cases superior to the straightforward VAR test. As a side-effect this procedure allows a distinction between the short-run and the long-run dynamics. As $b(L)$ describes only the short-run dynamics, $\alpha_1 = 0$ implies that di_t is not Granger-caused by dp_t in the long-run while there might be dependencies in the short-run. Note that the Engle-Granger representation theorem implies that at least one α_{1t} is non-zero in the bivariate system (8).

Granger and *Lee* (1989) propose the nonsymmetric error-correction model

$$(9)\quad \begin{pmatrix}\Delta di_t\\ \Delta dp_t\end{pmatrix} = \begin{pmatrix}\mu_1\\ \mu_2\end{pmatrix} + b(L)\begin{pmatrix}\Delta di_{t-1}\\ \Delta dp_{t-1}\end{pmatrix} + \begin{pmatrix}\alpha_1^+\\ \alpha_2^+\end{pmatrix}(di_{t-m} - \beta dp_{t-m})^+ + \begin{pmatrix}\alpha_1^-\\ \alpha_2^-\end{pmatrix}(di_{t-m} - \beta dp_{t-m})^- + \begin{pmatrix}\epsilon_{1,t}\\ \epsilon_{2,t}\end{pmatrix}$$

where $(di_t - \beta dp_t)^+ = \max(di_t - \beta dp_t, 0)$ and $(di_t - \beta dp_t)^- = \min(di_t - \beta dp_t, 0)$. This model allows for different adjustment coefficients for positive and negative deviations from the cointegration relationship. A higher absolute value of the coefficient for positive deviations would imply that the corresponding series responds more quickly to a positive equilibrium error compared to a negative one. Since the estimator for β is super-consistent, testing for $\alpha^+ = \alpha^-$ can be done by estimating β in the context of a linear model, inserting $\hat{\beta}$ into (9), and applying the usual Wald test. For positive and negative deviations to be defined, the mean of the cointegrating relationship must be identified. The problem relates to the constant μ in (9). If $\mu_i = \alpha_i \mu_0$ $(i = 1, 2)$, equation (9) can be rewritten as

$$(10)\quad \begin{pmatrix}\Delta di_t\\ \Delta dp_t\end{pmatrix} = b(L)\begin{pmatrix}\Delta di_{t-1}\\ \Delta dp_{t-1}\end{pmatrix} + \begin{pmatrix}\alpha_1^+\\ \alpha_2^+\end{pmatrix}(\mu_0 + di_{t-m} - \beta dp_{t-m})^+ + \begin{pmatrix}\alpha_1^-\\ \alpha_2^-\end{pmatrix}(\mu_0 + di_{t-m} - \beta dp_{t-m})^- + \begin{pmatrix}\epsilon_{1,t}\\ \epsilon_{2,t}\end{pmatrix}.$$

Only in this case the mean of the cointegration relationship is well defined and so are positive and negative deviations. Note that the rejection of (10) would imply that a linear trend component is present in inflation differentials and/or in interest rate differentials.[7]

4. Estimation and Results

Using monthly data from 1980:1 to 1994:9[8], in the present paper we investigate the 10-year government bond yield differential and the inflation differential vis-à-vis Germany for Austria, Belgium, Denmark, France, Italy, and the Netherlands.[9] Inflation is measured by the consumer price index, all data are taken in first differences of logs. With the exception of Austria, this selection of countries draws upon their membership in the EMS which was formed in 1979. Instead of joining the EMS, Austria opted for a rigorous pegging of the Austrian Schilling to the Deutsche Mark during the same

[7] Generally the constant term μ can be decomposed into two parts, μ_0, which represents the intercept in the cointegrating relationship, and the remainder, μ-μ_0, which determines a linear trend in di_t and dp_t (*Johansen*, 1991).

[8] For Denmark: 1983:6 to 1994:9. The estimation period ranges from 1981:6 to 1994:9 (for Denmark: 1984:10 to 1994:9). Data source: International Financial Statistics of the International Monetary Fund.

[9] Luxembourg has been excluded from consideration as it forms a currency union with Belgium.

period which we assume economically equivalent to being a formal EMS member country.

Figures 1-6 display strikingly close comovements between inflation differentials and government bond yield differentials between Germany and the other European countries which joined the EMS before 1981. Except for Italy, these countries are characterized by a comparatively high degree of integration as measured by the correlation of demand and supply disturbances with Germany (*Helmenstein* and *Url*, 1993). The graphs suggest the division of the countries into two groups. Belgium, Denmark, France, and Italy (Figures 1-4) faced high inflation differentials vis-à-vis Germany at the beginning of the eighties which entailed several devaluations of their currencies between 1980 and 1987. The reduction of the inflation differentials was accompanied by a parallel decrease of the interest rate differentials. Nevertheless, financial markets seem to have retained some suspicion about the stability of the EMS parities after 1987. This conclusion can be inferred from the observation that an increase of the inflation rate relative to Germany continued to induce an increase in the interest rate differentials. By contrast, a reversal in the inflation differentials towards Germany for the first time since 1980 paired with still positive (or at best zero) interest rate differentials led to a remarkable gap between real interest rates. Supposing that real interest parity holds, lower inflation rates in the neighboring countries to Germany imply that the German inflation rate is expected to return to the conventional level of inflation in the future, that is, the Bundesbank is expected to exert a comparable degree of monetary stability as before unification.

Austria and the Netherlands (Figures 5 and 6), which have maintained a stable exchange rate to Germany during the eighties, form the second group. Temporary inflation differentials seem to have exerted little impact on interest rate differentials. Figure 6 demonstrates that in the case of the Netherlands positive inflation differentials have always been quickly eliminated.

We test the null hypothesis of a unit root by Augmented Dickey-Fuller tests (ADF-tests) (*Dickey* and *Fuller*, 1981) with a constant but without a trend. The inclusion of a trend is economically implausible as this would imply continuously widening interest rate differentials. The outcome of this test may depend on the order of the approximating autoregression. Taking up the advice of *Ng* and *Perron* (1993), we select the lag length by a sequence of t-tests starting with an autoregression of order 4. Table 1 presents the results of ADF-tests for interest rate differentials and inflation differentials with Germany. For interest rate differentials the hypothesis of a unit root cannot be rejected at the 10% level for any country under study, while for inflation differentials this holds true for all countries except for Austria and

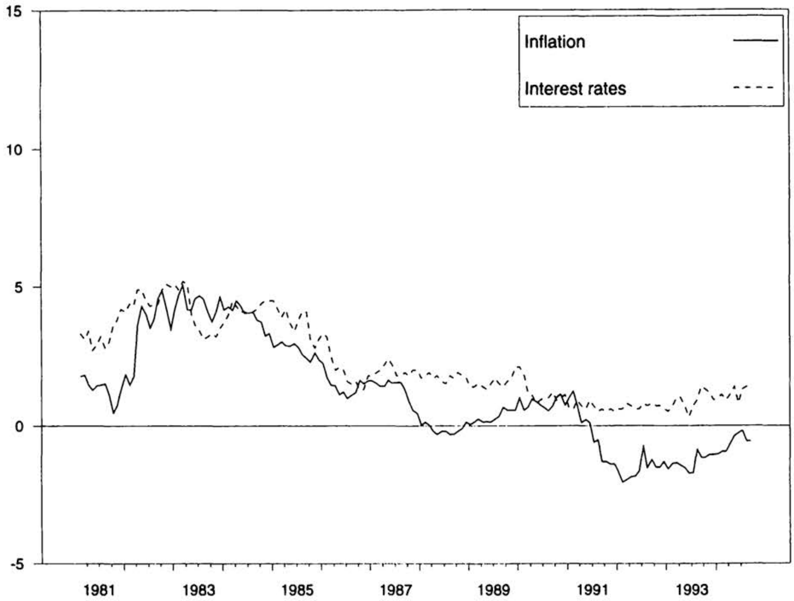

Figure 1: Belgium: Inflation and interest rate differentials

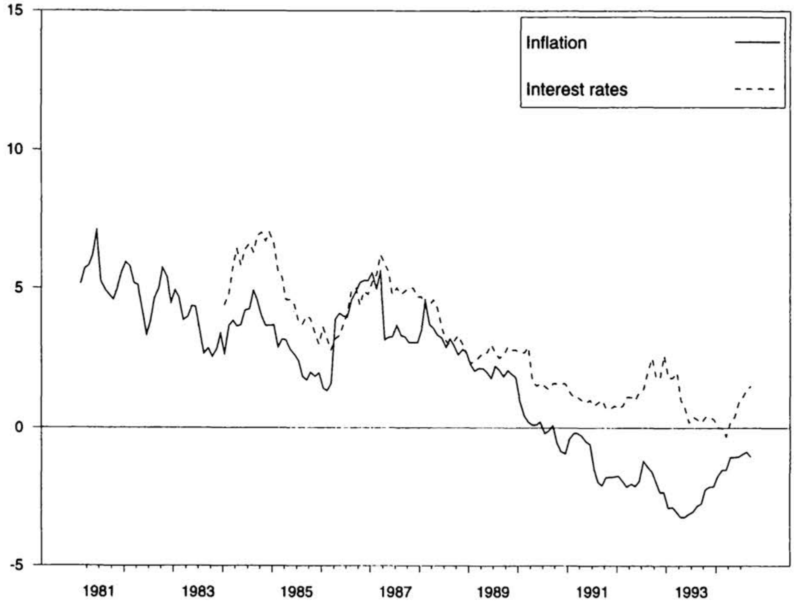

Figure 2: Denmark: Inflation and interest rate differentials

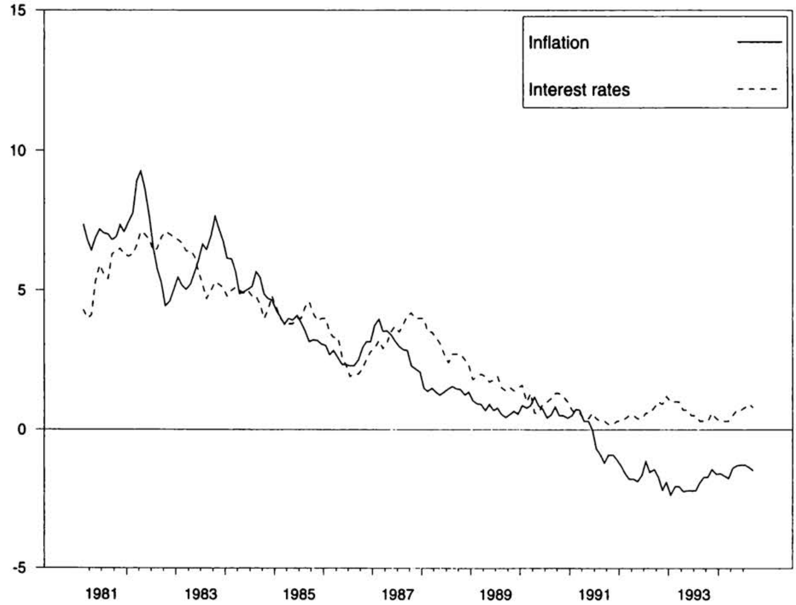

Figure 3: France: Inflation and interest rate differentials

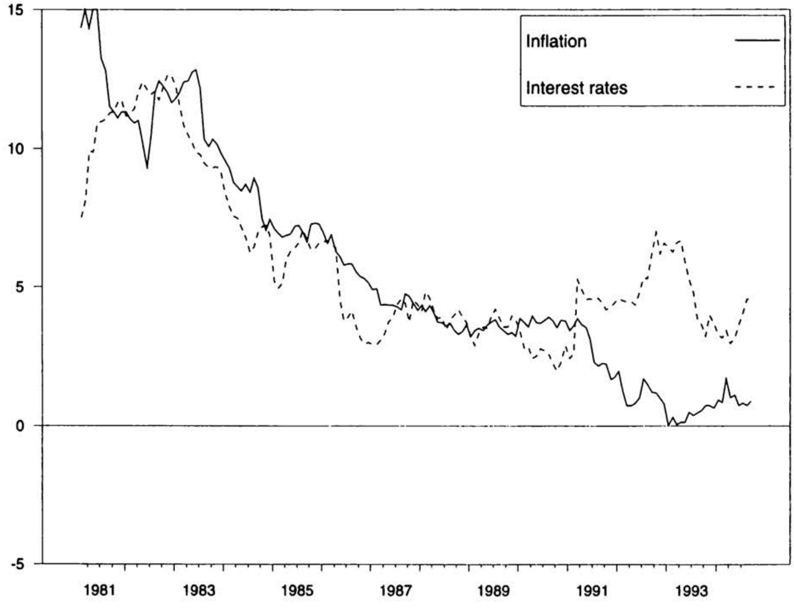

Figure 4: Italy: Inflation and interest rate differentials

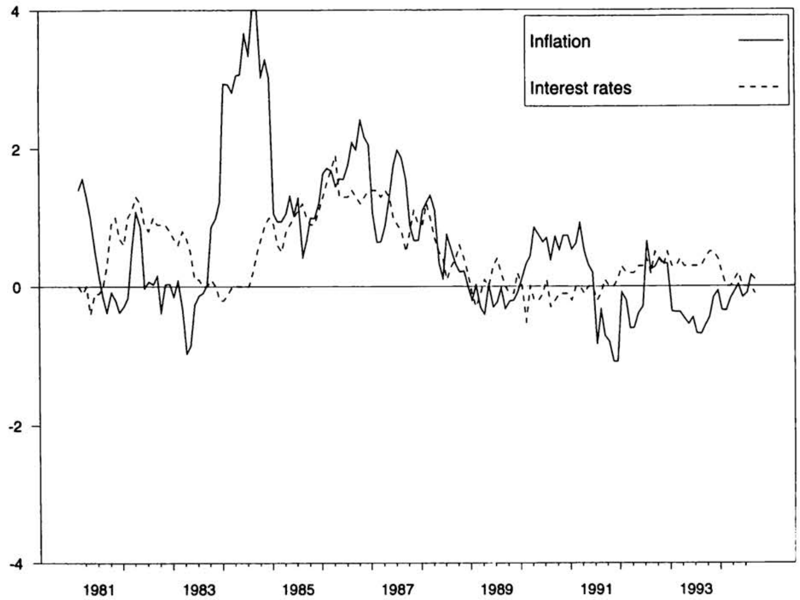

Figure 5: Austria: Inflation and interest rate differentials

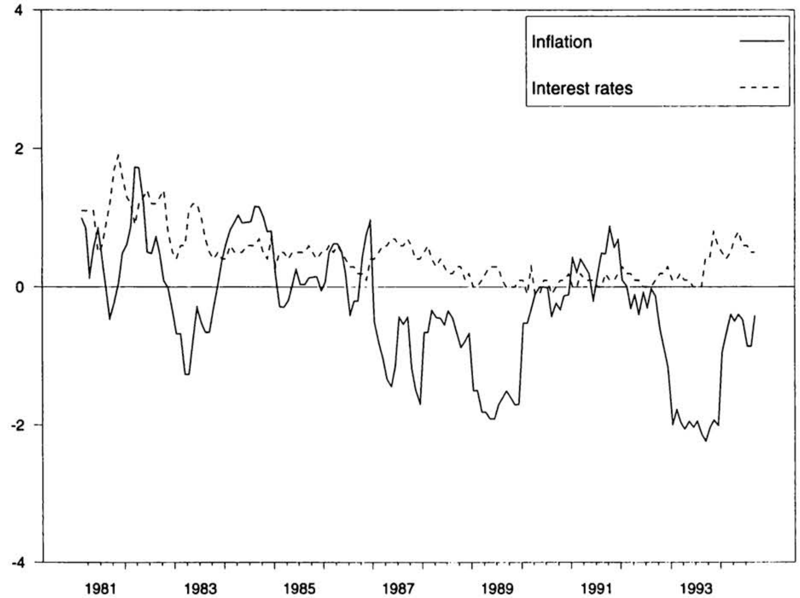

Figure 6: Netherlands: Inflation and interest rate differentials

the Netherlands. ADF-tests for the differenced series reject the hypothesis of a unit root at high significance levels so that all series can be regarded as first difference stationary.

Table 1

Unit Root and Engle-Granger Cointegration Tests

	Levels		First Differences		
Country	*di*	*dp*	Δdi	Δdp	Cointegration
Austria	-2.48 (0)	-2.85* (3)	-7.28*** (4)	-12.16*** (0)	-3.61** (0)
Belgium	-1.02 (2)	-1.02 (4)	-10.24*** (1)	-3.81*** (4)	-3.28** (0)
Denmark	-1.44 (0)	-1.47 (0)	-11.05*** (0)	-9.14*** (0)	-2.90 (0)
France	-1.02 (1)	-1.21 (1)	-11.34*** (1)	-9.22*** (0)	-3.48** (1)
Italy	-1.56 (1)	-2.44 (0)	-10.51*** (0)	-11.05*** (0)	-2.39 (1)
Netherlands	-2.40 (4)	-2.78* (3)	-9.94*** (4)	-8.09*** (2)	-3.09* (1)

1%, 5%, and 10% critical values of the ADF-tests: -3.43, -2.86, and -2.56;
for the cointegration tests: -3.90, -3.34, and -3.04 (*McKinnon*, 1991).
(.) represents the number of lags in ADF-tests;
*** significant at the 1 %-level;
** significant at the 5 %-level;
* significant at the 10 %-level.

Table 1 also shows Engle-Granger tests for cointegration using a constant but no trend. The tests confirm the existence of a cointegration relationship for all countries except for Denmark and Italy. As concerns Austria and the Netherlands, the unit root and cointegration tests do not provide clear-cut results. For both countries the ADF tests reject the nonstationarity of the inflation differentials at the 10% level but not the nonstationarity of the interest rate differentials. The fact that the null hypothesis of non-cointegration is rejected, however, implies that both series must be either stationary or integrated of order one. Statistical tests may provide asymmetric results in the sense that the failure to reject the null hypothesis may be due to a lack of power which does not necessarily imply the invalidity of the null. In our case the failure to reject the nonstationarity of interest rate differentials may be due to a lack of power of the ADF-test while the test statistic for the inflation differentials and the cointegration relationship are statistically significant at the 10%-level. We are therefore inclined to treat both series as stationary.

The above results are supported by the *Johansen* (1991) maximum likelihood estimates (Table 2). The lag length of the autoregressive polynomial was initially chosen by the Akaike information criterion but had to be increased for some countries due to autocorrelation of the residuals. The trace test for the number of cointegrating vectors indicates (at the 10% level) the existence of one cointegrating vector for Belgium and France. By contrast, in the case of Austria and the Netherlands two cointegrating vectors appear to be present, that is, both differentials are stationary. The test does not indicate the existence of a cointegration relationship for Denmark and Italy but for the former the trace statistic is close to the critical value at the 10% significance level. In the case of Denmark the failure to reject non-cointegration may be due to the shorter estimation period and some large deviations at the end of the sample (see Figure 2).

Table 2

Results of Johansen Cointegration Tests

Country	Lag length	Trace statistic #	β	Q (12)##	ARCH(4)##	LR_1 ###	LR_2 ###
Austria	1	14.20* 3.03*	-	12.49 16.52	0.41 2.25	-	-
Belgium	4	13.89* 0.42	0.65	8.23 4.49	1.16 0.96	8.29***	1.71
Denmark	1	12.67 0.11	0.75	8.16 12.09	0.14 1.77	3.52*	3.73*
France	2	21.89** 2.47	0.74	15.05 3.27	5.67** 3.83*	6.14**	8.63***
Italy	1	9.65 3.57*	-	7.09 10.88	0.47 0.05	-	-
Netherlands	4	14.14* 3.89*	-	9.97 14.20	0.92 0.85	-	-

\# For critical values of the trace statistics see *Johansen* and *Juselius* (1990).
\## The Ljung-Box statistic Q(12) and the LM test for ARCH effects are χ^2 distributed with 12 and 4 degrees of freedom, respectively. The first (second) value refers to the residuals of the inflation (interest rate) equation.
\### LR_1 and LR_2 are χ^2 distributed with 1 degree of freedom.
*** significant at the 1 %-level;
** significant at the 5 %-level;
* significant at the 10 %-level.

For Belgium, Denmark, and France Table 2 also provides the results of two different likelihood-ratio (LR) specification tests (*Johansen*, 1991). The first, LR_1, tests the null hypothesis H_0: $\beta = 1$, which follows from the theoretical relationship (6). This restriction is rejected for all three countries at different significance levels (Table 2). The second test, LR_2, tests for the existence of a linear time trend component, that is, H_0: $\mu_i = \alpha_i \mu_0$ $(i = 1, 2)$. This restriction is accepted for Belgium, whereas it is rejected for Denmark and France at the 10% level and at the 1% level, respectively. In the case of Belgium, the estimation of the model under the restriction $\mu_i = \alpha_i \mu_0$ $(i = 1, 2)$ essentially returns the same result. For Denmark, however, the trace statistics now indicates the existence of one cointegrating vector at the 10% significance level.[10] We therefore treat the Danish series as cointegrated.

Mutual Granger-causality tests according to the sequential method of *Toda* and *Philips* (1993, 1991) are presented in Tables 3 and 4. The first F-test (H_0: $\alpha_i = 0$) examines the significance of the error correction term, while the second F-test (H_0: $b_{ij}(L) = 0$) tests for the significance of the lag polynomial. The F-test for Granger-causality (H_0: $\alpha_i = b_{ij}(L) = 0$) examines whether both the lag polynomial and the adjustment coefficient of the error correction term are equal to zero.

For Belgium and France the tests indicate Granger-causality in both directions: interest rate differentials are Granger-caused by inflation differentials, and vice versa. Granger-causality is predominantly due to the significance of the adjustment coefficients while the coefficients of the lag polynomials appear to be insignificant. It is interesting to note that for shorter estimation periods (for example those starting in 1985) we find Granger-causality from inflation to interest rate differentials only. Though permitting a short estimation period only, the data for Denmark nevertheless show mutual Granger- causality. Granger-causality from inflation to interest rate differentials operates through the adjustment coefficient α_1. Conversely, interest rate differentials Granger-cause inflation differentials through the short-run dynamics.

Results presented for Austria and the Netherlands refer to the VAR (7) in levels due to the stationarity of the differentials. For Austria the tests reveal Granger-causality from inflation to interest rate differentials only while for the Netherlands we do not find Granger-causality in any direction. For Italy we tested for Granger-causality using a VAR in second differences but did not at all find evidence for Granger-causality.

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10 ARCH effects are generally found to be small. For some countries, however, the residuals in the inflation equation exhibit an excess kurtosis due to several outliers. This finding may be due to discretionary changes in taxation since the outliers mostly appear at the beginning of the year. Using dummy-variables to account for these did not change the results.

Table 3

Granger-Causality From Inflation Differentials to Interest Rate Differentials (F-tests)

Country	Number of lags	Error-correction terms H_0: $\alpha_1 = 0$	Short-run dynamics H_0: $b_{12}(L) = 0$	Both H_0: $\alpha_{12} = b_{12}(L) = 0$
Results for the VEC (8):				
Belgium				
(81:9-94:9)	4	3.41*	1.74	1.72
(85:1-94:9)		3.09*	1.38	1.37
Denmark	1	11.97***	1.10	6.34***
France				
(81:9-94:9)	2	7.91***	0.37	2.91**
(85:1-94:9)		5.11**	0.93	2.26*
Results for the VAR (7) in levels:				
Austria	1	n.a.	n.a.	4.04***
Netherlands	4	n.a.	n.a.	0.98

*** significant at the 1 %-level.
** significant at the 5 %-level.
* significant at the 10 %-level.

Table 4

Granger-Causality From Interest Rate Differentials to Inflation Differentials (F-tests)

Country	Number of lags	Error-correction terms H_0: $\alpha_2 = 0$	Short-run dynamics H_0: $b_{21}(L) = 0$	Both H_0: $\alpha_{21} = b_{21}(L) = 0$
Results for the VEC (8):				
Belgium				
(81:9-94:9)	4	7.01***	2.99**	3.30***
(85:1-94:9)		0.95	1.52	1.28
Denmark	1	0.04	5.37**	3.15**
France				
(81:9-94:9)	2	5.70**	0.94	3.46**
(85:1-94:9)		0.03	2.04	1.58
Results for the VAR (7) in levels:				
Austria	1	n.a.	n.a.	0.25
Netherlands	4	n.a.	n.a.	1.55

*** significant at the 1 %-level.
** significant at the 5 %-level.
* significant at the 10 %-level.

With Belgium and Denmark only two countries fulfill the preconditions for an application of the non-symmetric error-correction model (10). Wald tests for non-symmetric adjustment (H_0: $\alpha^+ = \alpha^-$), however, do not indicate asymmetries (Table 5).[11]

Table 5

Wald Tests for Asymmetries in the Response to Equilibrium Errors

Country	Interest rate differentials H_0: $\alpha_1^+ = \alpha_1^-$	Inflation differentials H_0: $\alpha_2^+ = \alpha_2^-$
Belgium	0.15	0.45
Denmark	0.66	2.01

The Wald statistic is χ^2 distributed with one degree of freedom.
*** significant at the 1 %-level;
** significant at the 5 %-level;
* significant at the 10 %-level.

5. Conclusions

The results of the foregoing analysis largely confirm the hypothesis that inflation differentials and interest rate differentials are indeed cointegrated. The relationship between inflation and interest rate differentials seems to be affected by the country-specific degree of credibility of monetary and fiscal policies. Recent literature stresses the importance of credibility in the impact of temporary imbalances in fiscal and monetary policies on the exchange rate, pointing out that the build-up of credibility is a long-lasting process. While it is difficult to quantify the concept of credibility, most observers will probably consent to an indicative ordering of EMS countries.

The Netherlands and Austria are well-known for a comparatively high degree of fiscal and monetary stability that is reflected by relatively sound fiscal balances, low inflation rates, and a lasting exchange rate peg to the DM that is also rationalized by their close trade links with Germany. By contrast, the other countries in the sample were more inflation-prone for the most part of the period under consideration. Belgium, Denmark, and Italy also carried high fiscal deficits during the early eighties. While Denmark and, to a minor extent, Belgium were able to reduce their net lending significantly during the eighties, public debt decreased in Denmark and stabilized in Belgium from 1987 on. In Italy, fiscal deficits have remained at

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11 The above results may not be correct if there is more than one cointegrating relation within the four-dimensional system of the inflation and interest rates of both countries. We checked for this possibility by applying the Johansen procedure to the four-dimensional system. Several authors pointed out (e.g. *Rünstler, Jumah,* and *Karbuz, 1995; Urbain, 1993)* that the efficiency of the Johansen method substantially declines with the dimensionality of the system which justifies the usage of the differentials themselves. In estimating the four-dimensional system for Denmark, the Netherlands, and Italy we find no cointegrating vector at all while there appears to be one for Austria, Belgium, and France. For Austria see also *Wörgötter* (1992).

high levels with public debt rising throughout. Taken together, the credibility of the Danish, French, Belgian, and Italian exchange rate peg would thus be considered weaker than the Austrian or Dutch.

We find that countries with highly credible policies have both stationary inflation and interest rate differentials with inflation Granger-causing interest rate differentials only for Austria. The returns of the Dutch and the Austrian hard-currency policy accrue in times of temporarily higher inflation rates vis-à-vis Germany (1991 and 1990 respectively) during which the interest rate differentials remain unaffected. The findings of the causality analysis back this conclusion. For EMS members with less credibility, by contrast, the respective differentials seem to be better characterized as non-stationary, with a cointegrating relationship arising from expected devaluations that are in line with cumulated inflation differentials. For Italy, the country with the least stable policies, a cointegrating relationship cannot be established, however. This may be due to a widely fluctuating risk premium component apart from inflation differentials.

The hypothesis of real interest parity was rejected for Belgium and France at the 5% level and for Denmark at the 10% level. In all three cases the coefficient in the cointegrating relationship, ß, was found to be smaller than one. It thus appears that interest rate differentials did not fully respond to inflation differentials or, in other words, real interest parity does not hold as stipulated by (6). Given that market participants expected devaluations to be equal to inflation differentials, nominal interest rate differentials did not fully offset expected devaluations. This finding suggests that within the EMS capital was not perfectly mobile during the estimation period.

Two more findings deserve further discussion. First, our estimates indicate that for Belgium, Denmark, and France from the mid-eighties on inflation differentials have Granger-caused inflation differentials but not vice versa while in the earlier period there was bidirectional Granger-causality. This change of the causality pattern coincides with the beginning of effective fiscal consolidation around 1984. Other factors than inflation differentials might therefore have played a major role in the formation of interest rate differentials, that is, large budget deficits may strongly influence the conduct of capital market participants.

Second, despite higher inflation rates in Germany than in most of the neighboring countries, interest rate differentials did not become negative in the early nineties. That is, Germany could draw on its „credibility capital“ during the phase of comparatively high inflation rates due to extensive government spending in the course of German unification. Applying a nonsymmetric error-correction model, however, we did not find evidence for an asymmetric response of interest rates to positive and negative deviations from the long-run equilibrium.

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Zusammenfassung

Die Theorie der ungedeckten Zinsparität sowie die Kaufkraftparitätentheorie implizieren das Bestehen einer langfristigen Gleichgewichtsbeziehung zwischen Inflations- und Zinsdifferentialen. In der vorliegenden Arbeit untersuchen wir die Trendeigenschaften dieser Differentiale für die EMS-Mitgliedsländer im Vergleich zu Deutschland. Den Ergebnissen von Einheitswurzeltests zufolge sind die niederländischen und österreichischen Differentiale jeweils stationär, während Kointegrationstests für Belgien und Frankreich die Existenz eines kointegrierenden Vektors anzeigen. Kausalitätstests zufolge verhalten sich Inflationsdifferentiale im allgemeinen Granger-kausal zu Zinssatzdifferentialen. Bei einzelnen Ländern läßt sich für die Zeit vor 1985 zusätzlich eine Granger-Kausalität von Zinsdifferentialen für Inflationsdifferentiale feststellen. Diese Resultate deuten darauf hin, daß der Glaubwürdigkeit der Geld- und Fiskalpolitik eine wesentliche Bedeutung für die Reaktion der Anleihemärkte auf die Preisdynamik zukommt.

Abstract

Uncovered interest parity and purchasing power parity suggest an equilibrium relationship between inflation and long-term interest rate differentials. In this paper we investigate the time trend properties of these differentials for the EMS members vis-à-vis Germany. The results of unit root tests indicate that for the Netherlands and for Austria both differentials are stationary while for Belgium and France we find a cointegration vector that represents the hypothesized relationship. Causality tests provide empirical evidence that inflation differentials generally Granger-cause interest rate differentials. For specific countries we also find Granger causality from interest rate to inflation differentials for the period before 1985. These results support the view that the credibility of monetary and fiscal policies plays an important role in determining the reaction of bond markets to inflation.

JEL Klassifikation: F15, C32

Keywords: Economic integration, cointegration, uncovered interest parity, European Monetary System, long-run causality